The impact of price competitiveness on UK producer price behaviour

Colin Ellis
and
Simon Price

Working Paper no. 178

Monetary Analysis Division, Bank of England, Threadneedle Street, London, EC2R 8AH.
Email: colin.ellis@bankofengland.co.uk, simon.price@bankofengland.co.uk

The views expressed are those of the authors and do not necessarily reflect those of the Bank of England. We are grateful for comments by Jenni Greenslade and Jan Groen.

Copies of working papers may be obtained from Publications Group, Bank of England, Threadneedle Street, London, EC2R 8AH; telephone 020 7601 4030, fax 020 7601 3298, e-mail mapublications@bankofengland.co.uk

Working papers are also available at www.bankofengland.co.uk/wp/index.html

The Bank of England's working paper series is externally refereed.

ISSN 1368-5562

Contents

Abstract

Modern open-economy macro models emphasise pricing-to-market behaviour. It is possible that domestic pricing behaviour might be affected by import (competitors') prices, and this is a commonly used variable in empirical work on pricing. But there is theoretical ambiguity and a potential identification problem. Cointegrating techniques are used in an attempt to resolve this, using the most appropriate data set (producer prices). Some evidence is found for the existence of two long-run relationships. The first of these is interpretable as a price mark-up or factor demand relationship, and competitors' prices can be excluded from it. The second equation can be interpreted as a long-run equilibrium price relationship equating domestic and foreign prices. This raises the possibility that single-equation estimates indicating a role for foreign prices in domestic price determination may mislead. However, the results are for producer prices and may not necessarily be extended to other indices.

Key words: Pricing, competitiveness.

JEL classification: D40, E30.

Summary

There are relatively few papers analysing the price mark-up equation. This is despite the fact that the role of price-setting in macroeconomics has come strongly to the fore recently. The 'new' Phillips curve is interpreted as a dynamic pricing equation, where marginal costs are proxied by the output gap, or, perhaps more satisfactorily, by unit labour costs. Within the literature, it has usually been taken as given that there should be a role for competitors' (import) prices. Yet there is theoretical ambiguity, and identification is a neglected issue. This is important for policy, as 'competitor' is synonymous with 'foreign' in this literature, and we know from the New Open-Economy Macro literature that pricing behaviour of importers is important when we consider the monetary transmission mechanism. Evidence from the existing literature using single-equation estimates does suggest such a relationship exists, and if this is the case, there are implications for the monetary transmission mechanism.

To help better understand the economic processes at work, we examine UK producer prices. This sector is a natural one to examine, because most output is tradable and the relevant economic model is likely to be appropriate. We relax the customary assumption of Cobb-Douglas production technology. There is a potential identification problem, as in principle there may be two long-run relationships – one that we will call a long-run price relationship (LRP: not necessarily purchasing power parity in the sense it is normally understood), and the other the optimal mark-up. Cointegrating techniques are used in an attempt to resolve this identification problem. We also have a proxy for competitiveness, which is intended to match comparable import and domestic prices. Some evidence is found for the existence of two separately identifiable long-run relationships. The first of these is interpretable as the price mark-up (or, equivalently, factor demand) relationship, and competitors' prices can be excluded from it. The second equation can be interpreted as a long-run equilibrium price relationship equating domestic and foreign prices.

This raises the possibility that single-equation estimates indicating a role for foreign prices in domestic price determination may unintentionally mislead. The results are for producer prices and may not necessarily be extended to other indices. But they suggest the possibility that the structural price mark-up equation for UK manufacturing does not depend upon foreign prices. This relationship appears to equilibrate via labour demand or, in terms of our modelled variables, productivity. However, there is evidence for a separate link between import and domestic producer prices, which might be thought of as the general equilibrium long-run relationship, which equilibrates through all three variables in our system. Thus, there is a suggestion that the reason why single-equation estimates find significant effects in price equations is that they conflate the structural and general equilibrium relationships.

1 Introduction

There is very little single-country analysis of the price mark-up equation. For the United Kingdom, papers include Martin (1997), Smith (2000) and Price (1991) and (1992). There is even less cross-country work.(1) This is despite the fact that the role of price setting in macroeconomics has come strongly to the fore recently.(2) The 'new' Phillips curve (eg, Galí and Gertler (1999)) is interpreted as a dynamic pricing equation, where marginal costs are proxied by the output gap, or, perhaps more satisfactorily, by unit labour costs.(3) Within the literature, it has usually been taken as given that there should be a role for competitor's (import) prices. Yet (as we argue below) identification is an important issue not addressed in these papers. Moreover, the theoretical justification for competitors' prices entering the price mark-up equation is not always clearly articulated; it is generally ambiguous with respect to the sign, and there may not be an effect at all. This is important for policy, as 'competitor' is synonymous with 'foreign' in this literature, and we know from the New Open Economy Macro literature that pricing behaviour of domestic producers and importers is important when we consider the monetary transmission mechanism. This pricing behaviour is often described as pricing-to-market (PTM),(4) and we shall use this as shorthand for models where competitors' prices affect firms' structural pricing decisions, although we will be careful to spell out precisely what this means. The literature developed in the analysis of importers' decisions, but it should be clear that the problem is symmetrical and applies equally to domestic producers.

In this paper we re-assess the results of previous work, paying particular attention to the identification problem. To help better understand the economic processes at work, we examine UK producer prices. This offers a more satisfactory sector than the obvious alternative, whole-economy prices, as the (profit maximising) theory is more likely to be applicable, and the goods in question are more likely to be traded. We also pay attention to identification problems not resolvable within a single-equation approach, and look at the implications of CES production technology. In Section 2 we spell out some theory. Section 3 examines some competitiveness series, while in Section 4 we briefly describe the data. In Section 5 we report the cointegration results, while Section 6 concludes. In a nutshell, we conclude that competitors' prices do not affect the long-run producer price mark-up.

---

(1) This disregards the large literature on purchasing power parity, which does not attempt, except indirectly, to estimate the mark-up over costs. Asteriou *et al* (2000) and Alogoskoufis *et al* (1990) are exceptions. In both papers, panels of about 20 OECD countries are used to estimate price mark-up equations. In the former manufacturing data is examined, using dynamic panel and panel cointegration techniques. Prices are found to depend pro-cyclically on capacity utilisation, and positively on competitors' prices. In the latter, the emphasis is explicitly on the role of competitors' prices.

(2) See the references cited in Martin (1997).

(3) This relies upon the maintained hypothesis that production technology is Cobb-Douglas. Relaxing this assumption allows a more general specification of the marginal cost.

(4) Marston (1990) defines pricing-to-market in a natural manner as the tendency for firms to choose to set different prices in different national markets because of differing market conditions.

2 Theory

The most basic economic theory tells us that the price an imperfectly competitive firm will optimally charge increases as the demand curve becomes more inelastic.[5] Thus anything that makes demand less elastic will increase the profit maximising mark-up. It has often been taken as given in the empirical literature that 'pricing to market' is synonymous with competitors' prices having some impact on a firm's price. But it should be clear that the mechanism must work through the influence of relative prices on the elasticity of demand. In this section we begin by briefly surveying the theoretical literature. We then consider a possible identification problem. Finally, we specify a particular equation to serve as an empirical vehicle.

2.1 Market structure

The message of the theoretical literature is that it is by no means obvious that imperfectly competitive firms will alter their mark-ups as foreign prices change, or if they do, in which direction it will be. Thus the only way to resolve this issue is by empirical analysis. The literature can be sorted into two types.[6]

First, the elasticity of demand is simply assumed to be a function of foreign prices, perhaps as a consequence of linear product demand curves. For example, Dornbusch (1987) presents a model of a Cournot equilibrium between domestic and foreign producers of an homogenous good sold in an oligopolistic domestic market, where the domestic price equilibrium is a weighted average of the domestic and exchange-rate-adjusted foreign marginal costs.[7] The elasticity of demand has two factors: the relative number of domestic to foreign firms, and the ratio of marginal cost to price of foreign suppliers; the effect of an exchange rate appreciation on domestic prices is unambiguously negative but may well be less than one-for-one. But different demand curves will lead to different results. In some recent New Open Economy Macro[8] models there are relative price effects on elasticities. These firms are modelled in a simple imperfectly competitive framework, and behaviour is properly microfounded. But the location of the assumption is effectively pushed back one step by specifying utility functions which generate the required demand curves. For example, in Bergin and Feenstra (2000) translog preferences are employed.[9] Thus in these cases the existence and sign of a relative price effect is essentially an assumption.

---

(5) See Section 2.3 below.
(6) We are grateful for Alisdair Scott's input to this discussion.
(7) Despite the apparent appeal to strategic interaction arguments, the results are driven by the assumption of a linear demand curve.
(8) An excellent survey of the current state of this burgeoning literature is in Obstfeld and Rogoff (2000).
(9) This assumption is motivated by a desire to explain macroeconomic features, including persistence. The authors argue some form of non-(log-)linearity in demand is required in order to generate realistic properties.

Second, models of strategic interaction draw from the industrial organisation literature for theories of oligopolistic interaction between domestic and foreign firms in the domestic economy. A standard reference is Bulow, Geanakopolos and Klemperer (1985), and there the effect can go either way. Froot and Klemperer (1989) present a model in which market share matters for future profits. Given a real exchange rate appreciation, firms will decide whether to raise margins (raising present profits) or lower margins (raising future share and future profits) depending on whether they think that the movement is temporary or permanent. Hence the elasticity of demand can be negative as well as positive, depending on agents' judgments of the nature of the exchange rate shock, their expectations of future paths, and their expectations of their competitors' actions in the (Nash) game. This theory is ambiguous about effects, therefore.

2.2 Identification

The evidence is that it is possible to find a role for world or import prices in single equations. All of Martin (1997), Price (1991) and Smith (2000) find comparable results, with a 2:1 weight in favour of costs relative to world prices. But we may not be estimating what we think we are. As a general equilibrium model property, it seems very likely UK and world prices are connected. For example, consider an appreciation in the exchange rate following an exogenous fall in the money supply; this example helps to emphasise the nominal nature of the shock. Assume there are no pricing-to-market effects and UK product demand is iso-elastic. Then sterling-denominated import prices will fall, reflecting lower foreign marginal costs. Domestic margins remain unchanged, but domestic market share will fall. Exports fall and import substitution occurs, so UK output and jobs suffer. In the labour market earnings fall so UK marginal costs decline. Capacity utilisation falls and margins may decline further. Thus domestic prices fall. In most models we are likely to work with nominal shocks have no long-run real effects. So the relative price of imports and domestic goods should be unaltered. These are general equilibrium effects, and there is no structural relationship in the pricing equation: world prices do not enter the price equation.[10] But single-equation regression analysis on aggregate data cannot distinguish an equilibrium property from the structural relationship. Thus in estimation we might mistakenly estimate a combination of a long-run model property and the mark-up equation.

To be specific, look at a situation where purchasing power parity (PPP) is a plausible hypothesis, such as a competitive market for a homogeneous product: wood screws, or memory chips, perhaps. Because the market is perfectly competitive, the law of one price holds. Thus we have the PPP condition,

$$p_i = p^*_i \qquad \textbf{(1)}$$

[10] Another possibility is that policy-makers target world inflation.

But it is nevertheless true that profit-maximising firms set prices as an optimal mark-up over marginal costs.

$$p_i = mc_i + \mu_i \qquad \textbf{(2)}$$

where *i* indexes the product, p_i is the domestic price, p_i* the foreign price in domestic currency, mc_i marginal cost and μ_i is the mark-up, actually zero in this case.(11) Firms do not set prices, of course: firm entry equilibrates the market and ensures **(2)** holds. It is easy to see that a regression of the form

$$p_i = \alpha_0 + \alpha_1 mc_i + \alpha_2 p*_i \qquad \textbf{(3)}$$

might yield misleading results as it will be an arbitrary statistical combination of **(1)** and **(2)**. The same could be true in aggregate. Fortunately, if the variables under analysis are non-stationary, there may be ways to address the issue.

2.3 Specification of the mark-up equations

We have argued that the role of relative prices is an empirical question. In order to answer it, we need sufficient structure to identify an empirical relationship. We can derive the pricing (which is simply inverted factor demand) equation from standard theory. Then monopolistically competitive firms face a demand curve for output of the following form:

$$Y^d = D(Z, P/P*) \qquad \textbf{(4)}$$

where *P/P** is the price relative to competitors' and *Z* is any other demand shifter. Given this demand curve, it is possible (but not necessary) that the elasticity Σ is also a function of these variables. Production is determined by a constant returns to scale production function

$$Y^s = F(NA, K) \qquad \textbf{(5)}$$

where *N* is employment, *A* is labour augmenting technical progress and *K* capital. The first-order condition can be written as

$$P = \frac{W / AF_1}{1 - 1/\Sigma(Z, P/P*)} \qquad \textbf{(6)}$$

(11) Here and below, lower case indicates logs.

where W is the nominal wage, W/AF_1 consequently defines marginal cost and $1\text{-}1/\Sigma\ (Z, P/P^*)$ determines the mark-up. Assume constant elasticity of substitution (CES) technology,

$$Y = \gamma[\delta K^{-\theta} + (1-\delta)(Ne^{at})^{-\theta}]^{-1/\theta} \qquad \textbf{(7)}$$

where the elasticity of substitution σ is given by $1/(1+\theta)$ and a_t indicates technical progress. In this case, one version of **(6)** is

$$p = \alpha_0 + w - (1/\sigma)(y-n) - (\sigma-1)/\sigma a_t + \alpha_1(p-p^*) + \alpha_2(y-y') \quad \textbf{(8)}$$

Here y' is trend or capacity output: that is, in line with much of the previous empirical work cited above, the demand shifter Z is proxied by capacity utilisation, $(y-y')$. As for competitors' prices but for different reasons, the sign of the effect is uncertain.(12) The coefficient on technical progress a_t may be positive or negative, depending on whether σ is less than or greater than one. The unrestricted version of this equation is

$$p = \beta_0 + \beta_1 w + \beta_2(y-n) + \beta_3 t + \beta_4(p-p^*) + \beta_5(y-y') \qquad \textbf{(9)}$$

In the equation, (the log of) technical augmenting progress a_t is simply modelled as a time trend, so that $\beta_3 = -(1+\beta_2)\varphi$ where φ is a positive scale factor. Thus the CES specification implies that $\beta_3 > (<) 0$ if $\beta_2 < (>) -1$. Normally, we would impose static homogeneity, $\beta_1 = 1$,

$$p = \beta_0 + w + \beta_2(y-n) + \beta_3 t + \beta_4(p-p^*) + \beta_5(y-y') \qquad \textbf{(10)}$$

With Cobb-Douglas technology σ is unity so $w-(1/\sigma)(y-n)$ may be replaced by unit labour costs. Recalling from **(8)** and **(9)** that $\beta_2 = -1/\sigma$,

$$p = \beta_0 + (w-y+n) + \beta_4(p-p^*) + \beta_5(y-y') \qquad \textbf{(11)}$$

(12) There are arguments suggesting the mark-up is countercyclical. Bils (1989) or more recently Ireland (1998) argue that firms use booms to attract new customers; Rotemberg and Saloner (1986) and Rotemberg and Woodford (1995) argue that collusive behaviour is less likely in booms, although their argument is restricted to exceptional price wars. Chevalier and Scharfstein (1996) have a model in which capital market imperfections lead to countercyclical pricing. Rotemberg and Woodford (1991) provide evidence of countercyclical mark-ups for the United States. Note that a positive coefficient on capacity utilisation may also indicate rising marginal cost. Bils' (1987) paper relies largely on estimated countercyclicality in marginal costs, following inflexible employment levels, to provide evidence for countercyclical mark-ups. The overall evidence, discussed in Layard, Nickell and Jackman (1991, pages 339-40), is mixed. Price (1991) and Smith (2000) report positive effects of capacity utilisation on UK data.

We have treated this problem as one where prices are being determined, but it is equally a factor demand or output supply decision. Thus the equivalent (labour) factor demand equation (conditioning on output) is

$$n = y - \sigma(w-p) - \sigma\alpha_0 + (\sigma-1)a_t - \sigma\alpha_1(p-p^*) - \sigma\alpha_2(y-y') \qquad \textbf{(12)}$$

With Cobb-Douglas technology this simplifies to

$$n = y - (w-p) - \alpha_0 - \alpha_1(p-p^*) - \alpha_2(y-y') \qquad \textbf{(13)}$$

Hence in the Cobb-Douglas specification unit labour costs can be used in the pricing equation and there is no exogenous trend productivity term. If this assumption is relaxed, unit labour costs need to be split into real wages and productivity, and the exogenous trend productivity term appears. There are cross-restrictions on these parameters.

As discussed in Section 2.2, as well as our general factor pricing equation **(9)**, a second relationship could hold;

$$p = \alpha_0 + p^* + \alpha_1 t \qquad \textbf{(14)}$$

One interpretation of this equation is PPP, but we argued above that it can be helpful to think in terms of a long-run open economy equilibrium relationship, so we shall refer to this as the long-run price relationship (LRP). Note that in any case PPP is weaker than the law of one price (LOP). The latter is consistent with perfectly competitive markets, which requires goods to be perfect substitutes. Where markets are less than perfectly contestable (perhaps because of transport or other fixed costs) firms may have market power and the LOP need not hold. Arguably, this is an uncontroversial assumption. It does not in itself imply the kind of PTM effects discussed above. It is plausible that in some markets there is room for market power to be held permanently, but a limit on relative prices is nevertheless determined in the long run because at sufficiently high domestic prices all markets are contestable, given a particular level of fixed costs. So again, the structural equation may not contain world prices, even though in the long run potential market entry and investment ensure domestic prices are linked to world prices. We include a time trend because of the secular decline in transport costs. Obstfeld and Rogoff (2000) show that for reasonable parameterisations of preferences, declining transport costs can have a big impact on relative demand for domestic and foreign goods (and thus explain the falling 'home trade bias puzzle'), and therefore relative prices.

In the rest of this paper we estimate the system comprising **(10)** (or equivalently **(12)**) and **(14)** using the Johansen technique, testing for the number of cointegrating relationships and the overidentifying restrictions. We also test the Cobb-Douglas structure and for weak exogeneity, which informs us about the causal structure of the model.

3 Measures of price competitiveness

The PTM theory predicts that the mark-up is related to the competitors' relative price. This might be described as price competitiveness. This in turn is closely related to the idea of the real exchange rate, and we must be careful to use the right terms here. Often, the term implicitly follows the 'Samuelson-Balassa' definition, which is the ratio of traded to non-traded prices.(13) The economic model here is one where traded goods obey purchasing power parity. In that case, a change in the nominal exchange rate cannot change competitiveness in the traded sector. But in our imperfectly competitive world, the more relevant concept is the ratio of domestic to foreign traded prices, or the terms of trade.

There are a variety of more or less relevant relative prices. For example, imports of capital goods do not directly compete on price with domestically produced consumer goods. Although the relative price will be relevant for agents' choice of what proportion of consumer and capital goods to purchase, the two goods are very obviously differentiated and so not in direct price competition. Ideally, we require a price competitiveness measure that captures competition between two substitutable products. One obvious possibility is the ratio of domestic manufacturing output prices to an import price index, as manufacturing produces largely traded goods. But imports also include non-manufactured goods. Consequently, this simple ratio is unlikely to solely capture competitiveness.

We consider the ratio of producer to three foreign prices. These are the price of imports to the United Kingdom, major 6 (M6) export prices and M6 producer output prices. The resulting competitiveness series for these three measures are shown in Charts 1 to 3 below. Naturally, nominal exchange rate fluctuations dominate the cycle, but in each case the series rise by roughly 50% over the 30-year sample. This would be a remarkable systemic decline in competitiveness. On the assumption that higher competitors' prices allow higher margins, this would imply a large secular profit squeeze. But the UK manufacturing profit share has risen slightly over this period (Chart 4). Similarly, the manufacturing rate of return has risen (Chart 5). It seems likely that all three measures suffer from compositional or other characteristics that make them poor comparators for UK domestic producer output prices.

(13) Balassa (1964), Samuelson (1964).

Chart 1
Competitiveness: UK import prices

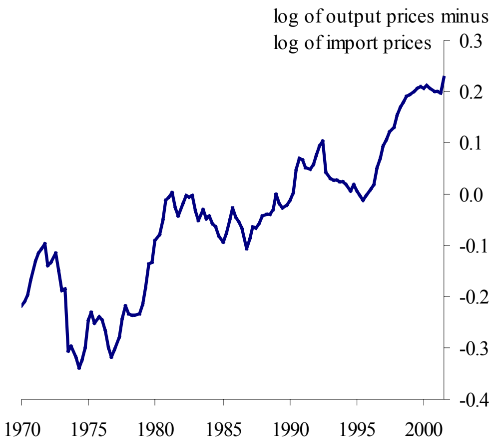


Chart 2
Competitiveness: M6 export prices

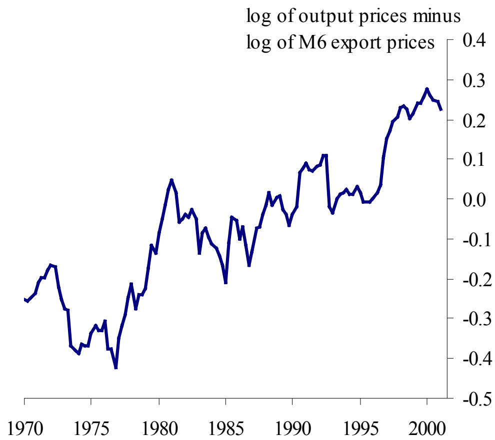


Chart 3
Competitiveness: M6 PPIs

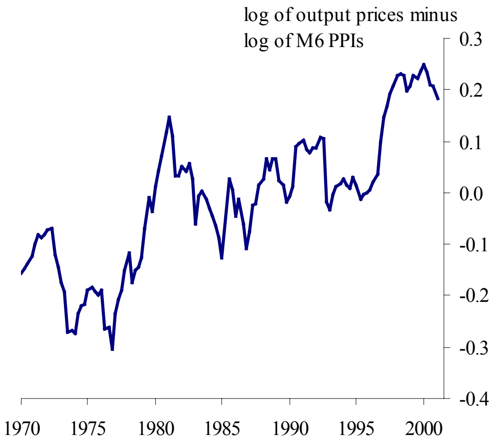


Chart 4
Manufacturing profit share[14]

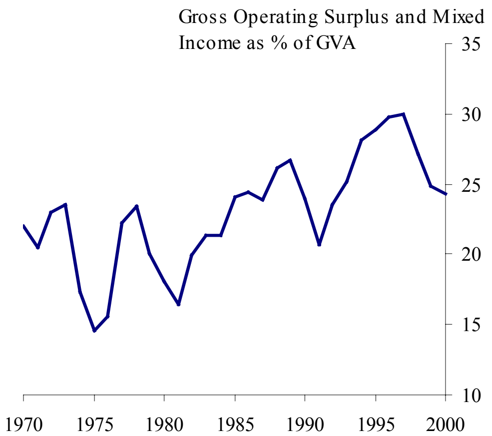
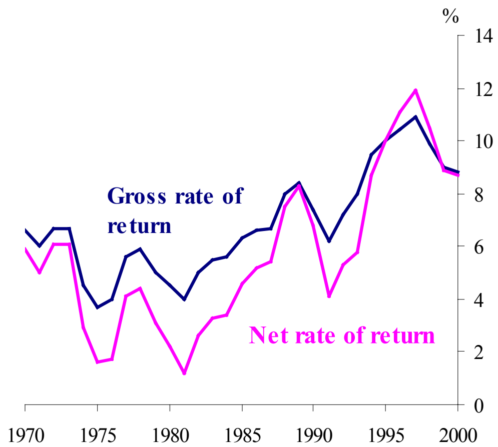


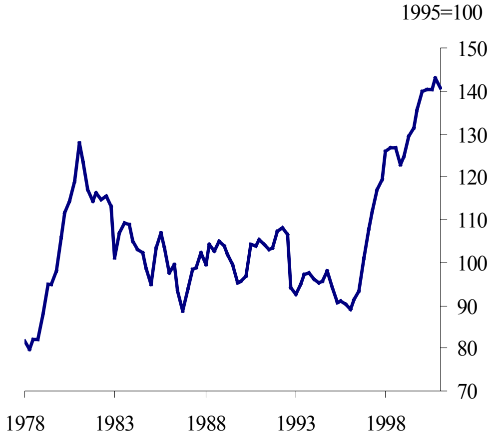
Chart 5
Manufacturing rate of return



(14) The numerator is for private rather than total manufacturing. For the total manufacturing profit share to have fallen, there would have to be a very substantial decline in the profit share of public manufacturers over the past 30 years. We could well believe this (eg due to privatisations), but the total manufacturing rate of return (Chart 5) has clearly not fallen. This indicates that (given capital) any fall in the profit share of public manufacturers has not resulted in a fall in the total manufacturing profit share.

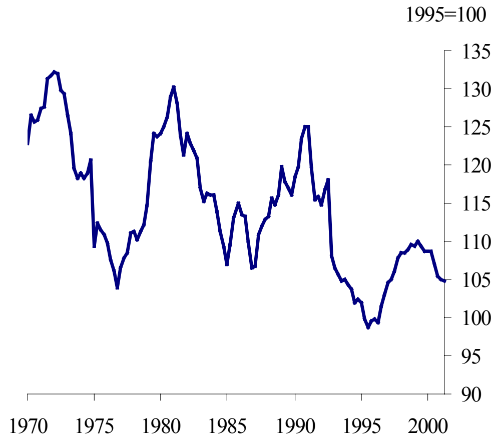
Alternative measures of competitiveness exist, however. Two that we could use are relative unit labour costs (published by the IMF) or import price competitiveness (ONS). These are shown in Charts 6 and 7 below.(15) The series closest to our conceptual model is the latter. It directly compares import and domestic prices for matched manufacturing goods at a disaggregated level. The chart shows that there has been a measured rise in competitiveness of about 20%, compared to the 50% fall suggested by Charts 1 to 3. It is this measure of price competitiveness that we use below. Unit labour costs clearly give an indication of relative costs, but we are interested in relative prices. Moreover, they may not be well measured, and are available over a shorter period.

Chart 6
Relative unit labour costs



N.B. A rise in the series indicates domestic unit labour costs becoming less competitive relative to those abroad.

Chart 7
UK import price competitiveness



N.B. A rise in the series indicates domestic prices becoming less competitive relative to imports.

(15) The import price competitiveness series has been spliced together from two separately published series to get a longer back run.

4 Data

The usual caveats about data limitations apply to our results. UK data for domestic manufacturing output, prices and wages are available on a quarterly basis back to the 1960s. Labour costs are calculated by taking into account employers' other labour costs.[16] The manufacturing price series is only available on a non-seasonally adjusted basis from the ONS. We seasonally adjusted with a standard X12 ARIMA process. To get a long-run employment series two separate series were spliced together: combining this with output yielded our productivity series. Our preferred competitiveness series (spliced together from two separately published series) is available from 1970, and our capacity utilisation series (CBI) is taken from the CBI *Quarterly Industrial Trends* survey, which dates back to January 1972. The four series we used after splicing, seasonal adjustment and labour cost adjustments are shown in the appendix.

Unit root tests suggest that the output price-wage ratio, productivity and competitiveness are all integrated of order 1 (see Table A), whether tested with or without a trend. The CBI capacity utilisation variable is marginally non-stationary with an ADF (1), but this result is not robust to different lags. A Phillips-Perron test cannot reject stationarity. Inspection of the data strongly suggests the series is stationary, and we proceed on this basis.

Table A: ADF tests statistics (1 lag)

	1 lag	1 lag and trend	4 lags	4 lags and trend
Level				
p-w	0.02	-2.67	-0.25	-2.40
y-n	-0.19	-1.68	-0.42	-1.89
p-p*	-2.00	-2.46	-1.89	-2.30
CBI	-2.46	-2.37	-2.73*	-2.80
First difference				
p-w	-9.25***	-9.20***	-5.35***	-5.32***
y-n	-7.39***	-7.35***	-4.98***	-4.95***
p-p*	-7.02***	-6.99***	-4.53***	-4.54***
CBI	-7.21***	-7.20***	-4.54***	-4.52***

*(**,***) indicates significant at 10% (5%, 1%) level

Samples (excl. lags): *p-w*, *y-n* and *p-p** 1970:1 2001:2, CBI 1971:4 2001:2

[16] We assume the proportion of wages in total costs is the same in manufacturing as the whole economy. We also experimented with a labour cost series that attempted to take into account the shift in the proportion of self-employed. Although individual coefficients changed slightly, the main results were unaffected.

5 Johansen results

Under standard notation, a vector error correction mechanism (VECM) can be written as:

$$\Delta X_t = \Gamma(L)\Delta X_{t-1} + \Pi X_{t-1} + \Phi D \quad \textbf{(17)}$$

where L is the lag operator, X is a matrix of I(1) variables, some of which may be weakly exogenous to the long-run relationship, and D is a set of I(0) variables both weakly exogenous to and insignificant in the long-run cointegration space. D may contain deterministic terms such as the constant and trend, and intervention dummies. All the long-run information is in the $(n\times n)$ Π matrix. Cointegration implies that this matrix contains a number $r < n$ of independent relationships, the long-run or cointegrating relationships. Π can always be decomposed into two matrices, $\alpha\beta'$, where both matrices are $(n\times r)$. Cointegration implies Π is of reduced rank, r, and so tests for cointegration are based on tests of the value of r. If $r = 0$ there is no cointegration. The α are the adjustment coefficients, also known as loadings, and the β are the long-run coefficients. The α also tell us about 'weak exogeneity'. A variable is weakly exogenous with respect to the long-run parameters in an identified cointegrating vector if the relevant element of α is zero. That is, if $\alpha_{ij} = 0$ (the loading on the jth cointegrating vector in the equation for the ith variable) then disequilibrium in that particular relationship is not equilibrated via the ith variable. This tells us something about 'long-run' causality, although there is nothing to prevent there being other short-run dynamic effects.

In our case we examine the set of I(1) variables $\{p\text{-}w,\ y\text{-}n,\ p\text{-}p^*\}$. We are testing for the number r of cointegrating relationships among this set, allowing for the I(0) variable y-y' and for a deterministic trend in the cointegrating space.

The method is maximum likelihood, and assumes Gaussian errors, so it is important to ensure that the residuals in the underlying VAR are normal and white noise. In order to determine the lag structure, we began by estimating an unrestricted VAR over the period 1971 Q4 to 2001 Q2 (the full data sample), adjusted for lags. Lag order selection criteria offered no definitive suggestion (see Table B),[17] but on the serial correlation criteria, eight lags were required. This number of lags implies that the VAR is almost certainly overparameterised, which reduces the power of the tests.[18] Thus we should err on the side of caution when determining the number of cointegrating vectors (use lower critical values). But the consequences of using too low a lag length are usually thought to be more severe. Hendry and Juselius (2000) conclude that

---

(17) But information criteria are not always considered to be helpful in determining lag lengths in cointegrating VECMs: Cheung and Lai (1993).

(18) There are many Monte Carlo studies of finite sample properties of the Johansen and other tests for cointegration, examining deviations from the maintained assumptions. Much of this literature is summarised in Maddala and Kim (1998). Results can be sensitive to the order in which assumptions about constants, trends and exogeneity are tested: Greenslade *et al* (2000).

‘[s]imulation studies have demonstrated that statistical inference is sensitive to the validity of some of the assumptions, such as, parameter non-constancy, serially-correlated residuals and residual skewness, while moderately robust to others, such as excess kurtosis (fat-tailed distributions) and residual heteroscedasticity.’ On normality, each equation in the VAR passed tests for skewness, but there was evidence for excess kurtosis at the 5% level in the *p-w* equation, and at the 1% level in the *y-n* equation.[19] While the kurtosis result should lead to caution in interpreting results, there is no reason to suppose inference is fatally flawed.

Table B: Lag length criteria

	Lag
LR test statistic	9
Final prediction error	9
AIC	10
SIC	1
H-QIC	2

Sample: 1973:4 2001:2

Table C: Johansen cointegration test

Null: r	Trace	Max-eigenvalue
=0	60.7***	35.7***
<=1	25.0*	19.8**
<=2	5.2	5.2

*(**,***) indicates rejection of null at 10% (5%,1%) level
Sample: 1973:1 2001:2

Using this lag structure, the Johansen cointegration tests indicated either one or two cointegrating relationships at the 5% significance level (Table C). However, two were selected at the 10% level. These tests are conducted without any exogenous variables. If we include the CBI terms as exogenous I(0) variables, there is still evidence for two vectors, with one test statistic suggesting two vectors at the 5% level. If tests conflict or are marginal, and given the comment above about the power of the tests, the appropriate procedure is to assume the larger number of vectors exist (Johansen (1995)). But in order to infer something about single-equation results, we first examine the results assuming one vector.

[19] We report estimates for a parsimonious system below. In that system we find that the deviation from normality is due to a residual outlier for one equation and not present for another.

5.1 Results assuming r = 1

Focusing on the long-run parameters in a VECM with a single long-run relationship,

$$\Pi X_{t-1} = \begin{bmatrix} \alpha_1 \\ \alpha_2 \\ \alpha_3 \end{bmatrix} \begin{bmatrix} \beta_{11} \beta_{12} \beta_{13} \end{bmatrix} \begin{bmatrix} x_1 \\ x_2 \\ x_3 \end{bmatrix}_{t-1} \quad \textbf{(18)}$$

Here x_1 is *p-w*, x_2 is *y-n* and x_3 is *p-p**. Normalising on β_{11} (the long-run coefficient on *p-w*), the cointegrating relationship and loadings are shown in Table D. We cannot reject the hypothesis that the coefficient on *p-p** is zero.(20)

An implication is that there is a unique cointegrating relationship between *p-w* and *y-n.* This appears to be the case on both the eigenvalue statistic at the 5% level and trace statistic at the 10% level. Results from estimating such a model are reported in Table E. However, the coefficient on productivity becomes large, implying a small elasticity of substitution, less than 0.2. We see below that in the original system *r* does indeed equal 2, and in the efficiently estimated system both vectors equilibrate via *p-p**. Technically, *p-p** is not weakly exogenous with respect to the parameters in the long-run relation of interest (see Chapter 8 in Johansen (1995)). The implication is that estimation of the restricted system is inefficient. We therefore move to the full system. We return to this issue in Section 5.4.

Table D: Unique cointegrating vector: including competitiveness

Sample: 1973:4 2001:2			
p-w	1.000000		
y-n	0.939943		
	[6.43267]		
*p-p**	0.002249		
	[0.01453]		
t	-0.000795		
	[-0.69513]		
Constant	3.813662		
Error Correction:	Δ (*p-w*)	Δ (*y-n*)	Δ (*p-p**)
Loading: vector 1	0.103295	-0.256124	0.097164
	[1.70024]	[-4.34786]	[1.02840]

t ratios in [brackets]

(20) The same results follow with the other measures of competitiveness described in Section 3. For the record, if we impose Cobb-Douglas technology (no time trend and a unit coefficient on productivity), the restriction is strongly rejected with a p-value of 0.00.

Table E: Unique cointegrating vector: excluding competitiveness

Sample: 1973:4 2001:2		
p-w	1.000000	
y-n	5.715834	
	[4.75405]	
t	-0.040142	
	[-4.18524]	
Constant	26.12147	
Error Correction:	Δ (*p-n*)	Δ (*y-n*)
Loading: vector 1	0.020604	-0.016796
	[4.09653]	[-3.13618]

But before we do so, there is the interesting question of what a single equation would have revealed. An auto regressive distributed lag (ARDL) specification parameterised as an error correction mechanism (ECM), restricted down from a general specification with four lags on all variables, reveals a significant error correction coefficient of –0.27 (*t* ratio 4.85), and long-run coefficients on *y-n* of –0.44 (*t* ratio 6.33) and *p-p** of –0.32 (*t* ratio 5.08), the latter result comparable with UK results reported in other papers. [(21)] This suggests a strong PTM effect. But in the light of the Johansen results reported above, with evidence for two cointegrating relationships, the single-equation results may be misleading. Interestingly, the results depend upon the competitiveness term. Exclude that, and the error correction coefficient falls to –0.07 and now has a *t* ratio of 1.49. This is insignificant at conventional levels, and furthermore under the null the regressors are non-stationary and the distribution is somewhere between asymptotic standard normal and Dickey-Fuller, so this will probably not reject the hypothesis of no cointegration, although we cannot be sure without simulating the model. The long-run coefficient on *y-n* is smaller (at –0.28) and insignificant (*t* ratio 1.04).

5.2 *Results assuming r = 2*

Given the results with a unique cointegrating vector, we proceed to the case where there are two. We estimate the long-run model

$$\Pi X_{t-1} = \begin{bmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \\ \alpha_{31} & \alpha_{32} \end{bmatrix} \begin{bmatrix} \beta_{11} & \beta_{12} & \beta_{13} \\ \beta_{21} & \beta_{22} & \beta_{23} \end{bmatrix} \begin{bmatrix} x_1 \\ x_2 \\ x_3 \end{bmatrix}_{t-1} \qquad \textbf{(19)}$$

To identify the cointegrating vectors we needed to apply restrictions to the two relationships. In the first relationship we identify the vector with a normalisation on *p-w* and setting $\beta_{13} = 0$ (competitiveness does not enter the equation). We refer to this as the factor demand (FD) equation. This restriction exactly identifies the equation, and is therefore not testing the model.

---

(21) Dummies were included to pick up 1973 oil crisis and 1979 crash outliers. There is no evidence of autocorrelation and the residuals are normal.

By contrast, in the second long-run relationship, we imposed the overidentifying and therefore testable restrictions that $\beta_{21} = \beta_{22} = 0$ and normalise on *p-p**, in line with **(14)**; as discussed in Section 2, we will refer to this as LRP. The second 'cointegrating relationship' implies *p-p** is trend stationary.[22] The model as a whole is overidentified. The long-run results and loadings are shown in Table F, with eight lags and eight lags of CBI included as an exogenous I(0) variable. The p-value for the restrictions is 0.89, easily accepted at standard significance levels. Apart from the loadings, the dynamics are not reported. Individual CBI terms are significant, but the sums are insignificantly different from zero in each equation; however, this is probably due to overparameterisation (see Table H below).

Table F: Two cointegrating relationships

Sample: 1973:4 2001:2
LR test for binding restrictions (rank = 2):

Chi-square(1)	0.018673		
Probability	0.891310		
	Coint Vector 1 FD	Coint Vector 2 LRP	
p-w	1.000000	0.000000	
y-n	0.931721	0.000000	
	[7.40561]		
*p-p**	0.000000	1.000000	
T	-0.000731	0.001562	
	[-0.71735]	[4.52450]	
Constant	3.786187	-4.832054	
Error Correction:	Δ (*p-w*)	Δ (*y-n*)	Δ (*p-p**)
Loading: vector 1	0.090660	-0.269344	-0.076492
FD(-1)	[1.18696]	[-3.63726]	[-0.67274]
Loading: vector 2	-0.012912	-0.010906	-0.177778
LRP(-1)	[-0.28483]	[-0.24815]	[-2.63443]

[22] In this system the relationship between the three series is driven by two cointegrating vectors, one of which describes the relationship between the I(1) series *p-w* and *y-n*, and the other which forces *p-p** to be stationary around a deterministic trend. In the short run, however, the three series are allowed to influence each other. So in order to test whether only the first two are I(1) and the third I(0) around a trend, implying an absence of a long-run influence between *p-w* and *y-n* on the one hand and *p-p** on the other, we followed two steps. First, we tested for the number of cointegrating vectors in the trivariate VECM and found that the system is driven by two cointegrating vectors, as implied by the current null. Conditional on these results we then tested whether *p-w* and *y-n* are I(1) and *p-p** I(0) as linear restrictions on the two cointegrating vectors in the normal way, as performed above.

Table G: Tests for weak exogeneity

	p-value
p-w **wrt vector 1**	0.57
y-n **wrt vector 1**	0.00
p-p* **wrt vector 1**	0.72
p-w **wrt vector 2**	0.94
y-n **wrt vector 2**	0.96
p-p* **wrt vector 2**	0.01
p-w **wrt vector 1, vector 2**	0.36

Sample: 1973:4 2001:2

There is an interesting pattern in the loadings. In particular, from inspection of the *t* ratios, *p-w* appears to be weakly exogenous to both relationships, implying *y-n* equilibrates factor demand (FD), and competitiveness equilibrates LRP. This is supported by the results in Table G.(23) But this VECM is almost surely overparameterised. Table H reports the results of estimating a more parsimonious system by the seemingly unrelated regression (SUR) method, maintaining the estimated long-run relationships but freeing the dynamics, and testing down. Dynamic variables were excluded where the t-statistic was less than 1.5(24) but the loading coefficients were not excluded even if they were insignificant. The number of dynamic coefficients were reduced from 96 to just 30; results are shown in Table H. In this case the hypothesis that *p-w* is weakly exogenous to the LRP relationship is rejected. But the loading on the factor demand relationship ('vector 1') is insignificant. So the long-run price mark-up equation does not equilibrate directly through prices. Instead, the adjustment is via labour demand, or product supply, if one prefers. Firms react to sub-equilibrium margins not by raising prices, but by lowering output and hiring less people. It is less easy to tell a story about the second relationship, LRP. Were this interpreted as PPP, however, then the mechanism is exactly right. Firms set prices equal to the world price, and then adjust factor inputs to equate prices to marginal cost.

This discussion of the causal mechanism based on the loadings is only suggestive. Recall that the notion of weak exogeneity is a statement about the exogeneity of a variable with respect to a set of parameters, in this case long-run. The VECM on which the estimates are based is a reduced form of the structural VAR; thus we cannot infer anything about the effect of (unidentified) structural shocks in the underlying system. The ratio *p-w* may not be – indeed, is unlikely to be – exogenous in the economic sense. But the important point is that although a long-run price mark-up or factor demand relationship exists, because the loading on *p-w* is zero, we would not expect to see a single-equation relationship if we estimate a single equation excluding competitiveness.

(23) Note that the test statistics and p-values are all for joint restrictions: not just the ones specified in the table, but also the overidentifying restrictions described above.
(24) Altering this to reject variables at normal significance levels had only a negligible impact on the results reported here. The broad test results reported later (ie acceptance/rejection of hypotheses) were unaffected.

Notice that the capacity utilisation terms are significant in each equation. This is particularly so in the productivity equation, where there is evidence of a positive rate of change effect.(25) The time trend in the first vector is insignificant, which is consistent with Cobb-Douglas technology. Furthermore, the coefficient on *y-n* is not significantly different from one, implying the elasticity of substitution is also around one. However, when we tested formally for Cobb-Douglas technology (replacing $w-(1/\sigma)(y-n)$ by unit labour costs and $\beta_2 = -1$), the restrictions were rejected at the 1% significance level.

(25) The rate of change restriction has a p-value of 0.12, accepted at normal significance levels.

Table H: Results for SUR system: *r* = 2

Estimation Method: Seemingly Unrelated Regression
Sample: 1973:1 2001:2

	Coefficient	Std. Error	t-Statistic	Prob.
Δ(p-w) equation				
FD(-1)*	**0.022125**	**0.041143**	**0.537761**	**0.5911**
LRP(-1)*	**-0.055499**	**0.025107**	**-2.210513**	**0.0278**
Δ((p-w)(-1))	-0.292275	0.079086	-3.695665	0.0003
Δ((p-w)(-3))	-0.166345	0.078496	-2.119139	0.0349
Δ((p-w)(-5))	-0.163505	0.077285	-2.115612	0.0352
Δ((p-w)(-6))	-0.229838	0.076471	-3.005544	0.0029
Δ((p-w)(-7))	-0.179077	0.079848	-2.242743	0.0256
Δ((p-w)(-8))	-0.275878	0.082785	-3.332469	0.0010
Δ((y-n)(-1))	-0.241483	0.069285	-3.485385	0.0006
Constant	-0.001793	0.004243	-0.422562	0.6729
CBI(-1)	-0.000736	0.000167	-4.402090	0.0000
CBI(-3)	0.000896	0.000211	4.243803	0.0000
CBI(-5)	-0.000455	0.000149	-3.041865	0.0026
R-squared	0.370011	Mean dependent var		-0.006500
Adjusted R-squared	0.295161	S.D. dependent var		0.013808
S.E. of regression	0.011592	Sum squared resid		0.013573
Durbin-Watson stat	2.113351			
Δ(y-n) equation				
FD(-1)*	**-0.201194**	**0.048695**	**-4.131737**	**0.0000**
LRP(-1)*	**0.015004**	**0.027147**	**0.552687**	**0.5809**
Δ((p-w)(-1))	0.432301	0.078194	5.528540	0.0000
Δ((p-w)(-3))	-0.188328	0.074027	-2.544051	0.0115
Δ((p-w)(-7))	0.177581	0.075340	2.357056	0.0191
Δ((p-w)(-8))	0.313482	0.077829	4.027818	0.0001
Δ((y-n)(-5))	0.129176	0.070312	1.837184	0.0672
Δ((y-n)(-6))	0.248306	0.069553	3.570040	0.0004
Δ((y-n)(-8))	0.211352	0.070057	3.016872	0.0028
Δ((p-p*)(-4))	-0.148830	0.049027	-3.035670	0.0026
Δ((p-p*)(-7))	-0.207660	0.050746	-4.092132	0.0001
Constant	0.000106	0.004735	0.022355	0.9822
CBI(-1)	0.001106	0.000196	5.655175	0.0000
CBI(-2)	-0.000646	0.000266	-2.425736	0.0159
CBI(-3)	-0.000622	0.000217	-2.862839	0.0045
CBI(-8)	0.000336	0.000115	2.916845	0.0038
R-squared	0.573591	Mean dependent var		0.006684
Adjusted R-squared	0.506263	S.D. dependent var		0.015826
S.E. of regression	0.011120	Sum squared resid		0.011748
Durbin-Watson stat	2.296959			
Δ(y-n) equation				
FD(-1)*	**-0.142020**	**0.060203**	**-2.359022**	**0.0190**
LRP(-1)*	**-0.178737**	**0.036242**	**-4.931796**	**0.0000**
Δ((p-w)(-3))	0.230208	0.110421	2.084824	0.0379
Δ((y-n)(-4))	0.259150	0.109208	2.372985	0.0183
Δ((p-p*)(-1))	0.176767	0.076149	2.321320	0.0209
Δ((p-p*)(-6))	0.215836	0.078285	2.757048	0.0062
Δ((p-p*)(-7))	0.289978	0.078203	3.708002	0.0002
Constant	-0.016785	0.006119	-2.743092	0.0065
CBI(-5)	0.001103	0.000316	3.489516	0.0006
CBI(-6)	-0.000713	0.000331	-2.153506	0.0321
R-squared	0.429024	Mean dependent var		-0.001664
Adjusted R-squared	0.379133	S.D. dependent var		0.021014
S.E. of regression	0.016558	Sum squared resid		0.028240
Durbin-Watson stat	1.932400			

* FD and LRP defined as in Table F.

To perform diagnostics, we re-estimated each of the three equations individually using OLS, which is consistent but inefficient. After accounting for a single outlying residual in the *p-w* and *p-p** equations, the three equations were all well-specified.[26] The residuals for all three pass tests for heteroscedasticity, autocorrelation, normality and ARCH. The conclusion appears to be that there are two long-run relationships, one of which comprises competitiveness and a time trend.

5.3 Results assuming fewer lags: robustness

Our results are based on a moderately high order VAR. As a robustness check, we re-estimated the model using less lags, namely 4 and 2. In the former case there was evidence of a single cointegrating vector at the 20% significance level, but not at the 10% level. When we estimated a model with a single cointegrating vector (CV), *p-p** could not be excluded from the CV and the loading on the *y-n* equation was insignificant. But the residuals from the unrestricted VECM with one CV are highly non-normal and fail tests badly (p-values below 0.01). Adding dummy variables models the non-normality, but leads to combinations of failures for heteroscedasticity, autocorrelation and ARCH. With two lags, once again there was only evidence for a single CV at the 20% significance level. When estimated assuming a unique relationship, the equations fail diagnostics as one might expect. Thus we are sure that the maintained lag structure is necessary.

5.4 Results assuming p-p is I(0): irreducibility*

We have determined that two cointegrating relationships exist within our set of variables. One of these implies *p-p** is effectively trend stationary,[27] which would normally lead to it being treated as an I(0) variable. Yet in this case we cannot simply include lags of *p-p** as I(0) variables, as the trend is necessary in the univariate model. As observed above, the price equation is only just identified and we cannot test the hypothesis of interest.

But given we believe *p-w* and *y-n* cointegrate, we should find such a set if we restrict attention to these two variables. This is an example of the notion of irreducible cointegration, formalised and explored by Davidson (1994, 1998). An irreducible cointegrating relation is one from which no variable can be omitted without loss of the cointegration property. Such relations may be structural or reduced form. The advantage of the procedure he suggests is that, under certain circumstances, when the model is overidentified, it enables the researcher to obtain information about the underlying structure directly from the data, and that is true in our case. The potential

---

[26] The dummy in the *p-w* equation was for 1974 Q1, the fuel crisis shock. The dummy in the *p-p** equation was for 1992 Q4, when sterling left the exchange rate mechanism (ERM).

[27] Oddly, the univariate tests in Table 1 rejected non-trend-stationarity. The Johansen test is evidently a more powerful unit root test in this dataset. Hansen and Juselius (2000, page 11) point out that a Johansen-based test differs from the Dickey-Fuller family of tests by testing stationarity as the null, conditional on reduced rank.

problem with the method is that of loss of efficiency when the cointegrating set is restricted, and that appears to be happening here. In effect, in Table E we test for an irreducible cointegrating relation; but we found the results were implausible. The results from the previous section imply that we need the information contained in the competitiveness series in order to efficiently estimate the pricing vector. But in order to pursue the Davidson approach to identification, we detrend *p-p** to create an I(0) exogenous series. In this case the set of I(1) variables is {*p-w, y-n*} and *r* can be at most 1. We condition on *y-y'* and detrended *p-p** as I(0) variables, using the long-run trend from an AR(8) for the detrending, in line with the lag length employed in the VECM. As mentioned in Section 5.1, there is evidence at the 10% (trace statistic) and 5% (max eigenvalue) level for a single cointegrating relationship. We include eight lags, as in the previous VECM estimates.

In contrast to the results in Table E, Table I shows that a long-run factor demand relationship can indeed be identified. The coefficient on *y-n* is again not significantly different from unity, and the time trend is insignificant. However, as in Section 5.2 formal tests of Cobb-Douglas technology were strongly rejected. Looking again at the loadings, *p-w* is weakly exogenous.(28) In the more parsimonious SUR estimates, reported in Table J, the significance rises to a p-value of 0.070. As we found in the SUR results that *p-p** is not weakly exogenous to the long-run parameters, the VECM assuming *r*=2 is to be preferred when estimating parameters. But we have been able to resolve the identification issue. Relative competitors' prices do not enter the expression determining the long-run mark-up.

Table I: Unique cointegrating vector: detrended *p-p I(0)**

Sample: 1973:4 2001:2

Cointegrating Eq:	Coint vector 1 FD	
p-w	1.000000	
y-n	0.875336	
	[6.63873]	
t	-6.76E-05	
	[-0.06730]	
Constant	3.509015	
Error Correction:	D(G)	D(H)
Loading: vector 1 FD(-1)	0.097973 [1.40295]	-0.283018 [-4.17224]

(28) It may be that this follows from the definition of the dependent variable as $\Delta(p-w)$, but the result holds when we insert it into a system where the dependent variable is defined as Δp.

Table J: Results for system: *p-p** I(0)

Estimation Method: Seemingly Unrelated Regression
Sample: 1973:4 2001:2

	Coefficient	Std. Error	t-Statistic	Prob.
Δ(p-w) equation				
FD(-1)*	**0.097969**	**0.053755**	**1.822515**	**0.0700**
Δ((p-w)(-1))	-0.391337	0.086413	-4.528684	0.0000
Δ((p-w)(-2))	-0.177820	0.090526	-1.964302	0.0510
Δ((p-w)(-3))	-0.364481	0.082973	-4.392772	0.0000
Δ((p-w)(-4))	-0.131341	0.078932	-1.663988	0.0978
Δ((p-w)(-5))	-0.240458	0.081170	-2.962389	0.0034
Δ((p-w)(-6))	-0.211855	0.076239	-2.778843	0.0060
Δ((p-w)(-7))	-0.171412	0.077245	-2.219083	0.0277
Δ((p-w)(-8))	-0.287532	0.079873	-3.599875	0.0004
Δ((y-n)(-8))	-0.273865	0.067412	-4.062588	0.0001
Constant	0.004814	0.095199	0.050567	0.9597
CBI(-1)	-0.000772	0.000173	-4.468536	0.0000
CBI(-3)	0.000804	0.000206	3.904562	0.0001
CBI(-5)	-0.000488	0.000207	-2.355063	0.0195
CBI(-7)	0.000513	0.000239	2.145253	0.0332
CBI(-8)	-0.000434	0.000197	-2.200543	0.0290
(p-p*)(-2)	-0.059674	0.029992	-1.989652	0.0481
(p-p*)(-8)	0.058079	0.024424	2.377927	0.0184
R-squared	0.453263	Mean dependent var		-0.006310
Adjusted R-squared	0.353321	S.D. dependent var		0.013799
S.E. of regression	0.011097	Sum squared resid		0.011452
Durbin-Watson stat	2.212350			
Δ(y-n) equation				
FD(-1)*	**-0.298823**	**0.046459**	**-6.431895**	**0.0000**
Δ((p-w)(-1))	0.462075	0.081463	5.672195	0.0000
Δ((p-w)(-7))	0.157338	0.078541	2.003268	0.0466
Δ((p-w)(-8))	0.363644	0.080869	4.496700	0.0000
Δ((y-n)(-5))	0.118159	0.071028	1.663559	0.0979
Δ((y-n)(-6))	0.192158	0.072194	2.661702	0.0084
Δ((y-n)(-8))	0.193944	0.070966	2.732924	0.0069
Constant	0.035639	0.099824	0.357012	0.7215
CBI(-1)	0.001075	0.000210	5.119301	0.0000
CBI(-2)	-0.000602	0.000271	-2.226021	0.0272
CBI(-3)	-0.000516	0.000229	-2.252339	0.0255
CBI(-8)	0.000333	0.000114	2.921447	0.0039
(p-p*)(-3)	-0.082154	0.031410	-2.615545	0.0096
(p-p*)(-7)	-0.147288	0.054056	-2.724738	0.0070
(p-p*)(-8)	0.221296	0.052577	4.208991	0.0000
R-squared	0.559305	Mean dependent var		0.006684
Adjusted R-squared	0.495037	S.D. dependent var		0.015826
S.E. of regression	0.011246	Sum squared resid		0.012142
Durbin-Watson stat	2.223494			

* FD defined as in Table I.

6 Conclusions

There are theoretical reasons to suppose that producer prices might be affected by (foreign) competitors' prices. Single-equation estimates do suggest such a relationship exists, and if this is the case, there are implications for the monetary transmission mechanism. For example, an appreciation in the pound will have a permanent effect on prices in the United Kingdom: the exchange rate is not neutral. But there is a question of identification. In this paper, we have used cointegration techniques to examine UK producer prices. This sector is a natural one to examine, because most output is tradable and the relevant economic model is likely to be appropriate. We also have a proxy for competitiveness which is intended to match comparable import and domestic prices. Cointegration analysis is well known to be sensitive to assumptions, such as exogeneity and lag length, so our results must be treated with some caution. But they suggest the possibility that the structural price mark-up equation for UK manufacturing does not depend upon foreign prices. This relationship appears to equilibrate via labour demand or, in terms of our modelled variables, productivity. However, there is evidence for a separate link between import and domestic producer prices, which might be thought of as the general equilibrium long-run relationship, that equilibrates through all three variables in our system. Thus it may be that the reason single-equation estimates find significant effects in price equation is that they conflate the structural and general equilibrium relationships.

Appendix: The data series

The ratio of prices to labour costs, productivity, competitiveness and the CBI capacity utilisation balance are shown in the charts below. All variables apart from CBI are in log levels.

Chart A1: Ratio of price to labour costs

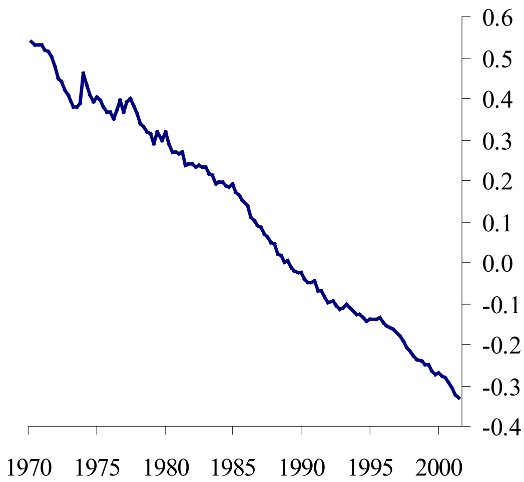


Chart A2: Productivity

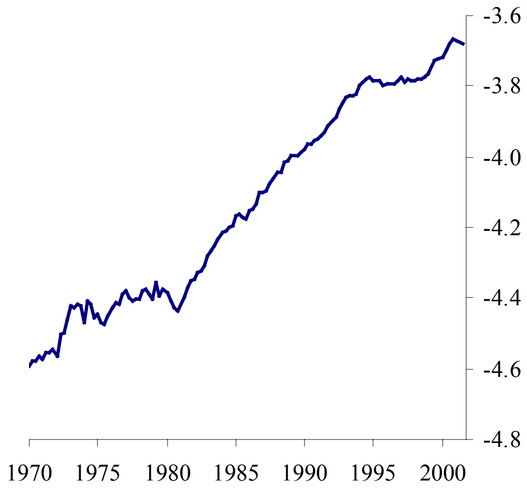


Chart A3: Competitiveness

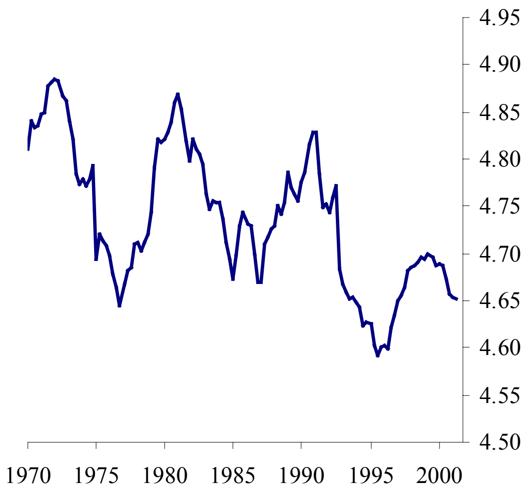
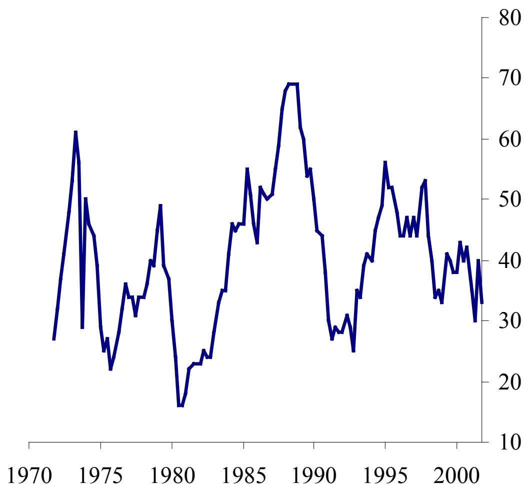


Chart A4: CBI capacity utilisation



References

Alogoskoufis, G, Martin, C and Pittis, N (1990), 'Pricing and product market structure in open economies: an empirical test', *CEPR Discussion Paper*, No. 486.

Asteriou, D, Lukacs, P, Pain, N and Price, S (2000), 'Manufacturing price determination in OECD countries; markups, demand and uncertainty in a dynamic heterogeneous panel', *City University Economics Working Paper*.

Balassa, B (1964), 'The purchasing power parity doctrine: a reappraisal', *Journal of Political Economy*, Vol. 72, page 496.

Bergin, P and Feenstra, R C (2000), 'Staggered price setting, translog preferences, and endogenous persistence', *Journal of Monetary Economic*s, pages 657-80.

Bils, M (1987), 'The cyclical behavior of marginal cost and price', *American Economic Review*, Vol. 77, pages 838-55.

Bils, M (1989), 'Pricing in a customer market', *Quarterly Journal of Economics*, Vol. 104, pages 699-718.

Bulow, J J, Geanakopolos, J D and Klemperer, P D (1985), 'Multimarket oligopoly: strategic substitutes and complements', *Journal of Political Economy*, Vol. 93, pages 488-511.

Cheung, Y and Lai, K (1993), 'Finite sample sizes of Johansen' likelihood ratio tests for cointegration', *Oxford Bulletin of Economics and Statistics*, Vol. 55, pages 313-28.

Chevalier, J A and Scharfstein, D S (1996), 'Capital-market imperfections and countercyclical markups: theory and evidence', *American Economic Review*, Vol. 86, pages 703-25.

Davidson, J (1994), 'Identifying cointegrating regressions by the rank condition', *Oxford Bulletin of Economics and Statistics*, Vol. 56, pages 103-08.

Davidson, J (1998), 'Structural relations, cointegration and identification: some simple results and their application', *Journal of Econometrics*, Vol. 87, pages 87-113.

Dornbusch, R (1987), 'Exchange rates and prices', *American Economic Review*, Vol. 77, pages 93-106.

Froot, K and Klemperer, P (1989), 'Exchange rate pass-through when market share matters', *American Economic Review*, Vol. 79, pages 637-54.

Galí, J and Gertler, M (1999), 'Inflation dynamics: a structural econometric model', *Journal of Monetary Economy*, Vol. 44, pages 195-222.

Greenslade, J V, Hall, S G and Henry, S G B (2000), 'On the identification of cointegrated systems in small samples: practical procedures with an application to UK wages and prices', *Centre for International Macroeconomics Discussion Paper*, No. 2000.04.

Hansen, H and Juselius, K (2000), 'CATS in RATS' (Evanston: Estima).

Hendry, D F and Juselius, K (2000), *Explaining cointegration analysis: part II*, unpublished.

Ireland, P (1998), 'Customer flows, countercyclical markups and the output effects of technology shocks', *Journal of Macroeconomics*, Vol. 20, pages 649-64.

Johansen, S (1995), *Likelihood-based inference in cointegrating vector auto-regressive models* (Cambridge: CUP).

Layard, R, Nickell, S and Jackman, R (1991), *Unemployment: macroeconomic performance and the labour market* (Oxford: Oxford University Press).

Maddala, G S and Kim, I-M (1998), *Unit roots, cointegration and structural change* (Cambridge: CUP).

Marston, R C (1990), 'Pricing-to-market in Japanese manufacturing', *Journal of International Economics*, Vol. 29, pages 217-36.

Martin, C (1997), 'Price formation in an open economy: theory and evidence for the United Kingdom, 1951-1991', *Economic Journal*, Vol. 107, pages 1,391-404.

Obstfeld, M and Rogoff, K (2000), 'The six major puzzles in international macroeconomics: Is there a common cause?', *NBER Working Paper*, No. 7777.

Price, S (1991), 'Costs, prices and profitability in UK manufacturing', *Applied Economics*, Vol. 23, pages 839-50.

Price, S (1992), 'Forward looking price setting in UK manufacturing', *Economic Journal*, Vol. 102, pages 497-505.

Rotemberg, J J and Saloner, G (1986), 'A supergame-theoretic model of price wars during booms', *American Economic Review*, Vol. 76, pages 390-407.

Rotemberg, J J and Woodford, M (1991), 'Markups and the business cycle', in Blanchard, O J and Fischer, S (eds), *NBER Macroeconomics Annual*, (NBER), pages 63-128.

Rotemberg, J J and Woodford, M (1995), 'Dynamic general equilibrium models with imperfectly competitive product markets', in Cooley, T F (ed), *Frontiers of business cycle research*, (Princeton University Press), pages 243-93.

Samuelson, P (1964), 'Theoretical notes on trade problems', *Review of Economics and Statistics*, Vol. 23, pages l-60.

Smith, P N (2000), 'Output price determination and the business cycle', *Economic Modelling*, Vol. 17, pages 49-69.